# LIBERALIZATION AND STOCK MARKET CO-MOVEMENT BETWEEN EMERGING MARKETS

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## Abstract

In this paper, we investigate the impact of trade and financial liberalization on the degree of stock market co-movement among emerging economies. Using a sample of 25 developing countries observed over 15 years, we estimate the impact of reforms which aim at opening these countries to trade and financial channels to the rest of the world. The estimation of time-varying cross-country correlations allows the econometric investigation to be performed using a panel data framework, raising hence the quality of the statistical inference. Our results offer strong support in favor of a positive impact of trade and financial liberalization reforms on the degree of cross-country stock market linkages.

JEL Code: F15, F36, G15, O10.

Keywords: stock market synchronization, financial integration, trade integration, emerging markets.

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# 1 Introduction

The extent of financial market synchronization is of the utmost importance for a large number of economic agents. The size and the evolution of the correlation between returns in international equity markets are crucial for appropriate portfolio selection. A good understanding of the determinants of stock market co-movements also sheds light on important issues like understanding the so-called home country bias puzzle (Lewis, 1999), i.e. the fact that international portfolios are insufficiently diversified and that investors tend to invest more on their own domestic markets.

Not surprisingly, the measurement of the cross-country linkages has been studied by some important papers (Bekaert and Harvey, 1995; Forbes and Rigobon, 1999). Meanwhile, the investigation of the determinants of cross-country financial interdependence has been studied by a large empirical literature aiming at identifying the role of a set of factors, such as trade intensity (Chinn and Forbes, 2004), financial development (Dellas and Hess, 2005), business cycle synchronization (Walti, 2005) and geographical variables (Flavin et al., 2002). Generally, these studies find some support for an explanatory power of these factors, although the results and conclusions differ significantly across studies. This divergence might be partly explained by the high degree of heterogeneity in the empirical approaches adopted by the literature. This heterogeneity concerns the sample of included countries (developed vs developing countries), the nature of the econometric approach (cross-section vs time-series), the measurement of market co-movement and, last but not least, the nature, as well as the measurement of explanatory factors.

Another important part of the literature has been concerned by the developments in terms of financial liberalization. There is a strong evidence that over the last twenty-five years, financial markets have become more integrated and that investors are able now to invest both in domestic and foreign assets. Obviously, the process of financial globalization has not followed a linear time trend during the 20th century. Obstfeld and Taylor (2005) show for instance that capital market integration follows a U-shaped pattern, with the interwar period exhibiting a relatively high degree of capital controls. Nevertheless, since the end of the so-called Bretton-Wood period (1944-1973), developed and developing countries have enjoyed a new wave of financial openness. While the exact timing of financial liberalization remains somewhat controversial on a country by country basis (see Bekaert et al., 2003), there is a broad consensus among economists to claim that capital markets are much more integrated today than 30 years ago.

The striking liberalization of financial markets has given rise to a large set of studies investigating the consequences of such a process. For instance, Bekaert et al. (2001, 2005) investigate the impact of financial liberalization on economic growth and investment in developing countries. Bekaert and Harvey (2000) study the impact of capital market integration for stock market correlations using a set of case studies while Quinn and Voth (2006) look at the same impact over a very long period (more than 100 years) for developed countries. Eizaguirre and Biscarri (2006) look at the effects of financial liberalization of emerging markets in terms of volatility.

In this paper, we revisit the empirical investigation of the relationship between cross-country stock market correlations with trade and financial liberalization. We depart from the existing literature in several aspects. First, we focus on the financial interdependence between emerging economies, avoiding mixing up developed and developing countries.[1] Compared to mature mar-

[1] A noticeable exception is provided by Pretorius (2002). As emphasized in her paper, previous studies tend to

kets, emerging markets are known to exhibit contrasting features such as market liquidity, agent's supervision and access to international capital markets. A separate investigation allows the identification of the determinants specific to emerging markets. A couple of these countries such as Thailand of Malaysia have even considered to reduce the exposition of the foreign markets, which suggest that there is some doubt on the gains drawn from liberalization.

Second and importantly, in contrast to a major part of the literature (Dellas and Hess, 2005 for instance), we combine the use of cross-section and time series data, explicitly controlling for unobserved heterogeneity both in the cross section and the time series dimensions. This approach is made possible by the estimation of a time-varying measure of correlation based on the use of realized moments. This method developed in the area of financial econometrics (Anderson and Bollerslev, 1998) makes use of the high frequency of financial data and increases the quality of the estimation at lower frequencies. In turn, the combination of cross-section and time-series data allows some light to be shed on the role of time-varying factors (such as liberalization episodes or trade intensity). Accounting for unobserved heterogeneity is important here as it allows unobservable factors, that explain the variation in stock market synchronization, to be controlled. In particular, inclusion of time dummies accounts for the impact of common shocks, such as major financial crises with wide contagion effects, that affects all pairs of countries. The need to control for these common shocks when explaining correlations has been stressed by several authors, including Chinn and Forbes (2004). Likewise, the presence of country-pairs effects accounts for the effect of factors such as geographical distances or differences in industrial specialization that are constant over time or that display a lot of inertia. The inclusion of these time and individual effects leads to a parsimonious specification and allows us to concentrate on specific factors such as financial and trade liberalization.

Third, the use of a dynamic panel data framework allows the possible dynamics of the correlation to be taken into account. In most, if not all, the literature to date, correlation is supposed to be constant over time, and so fluctuations, due for example to business cycles, are neglected. Tests are used to justify the dynamic dimension of the panel leading to a distinction between the short and the long run effect of the variables on the stock market correlations.

Fourth, the focus on developing countries and the use of panel data allows us to study the specific role of financial and trade liberalization. These variables might explain a major part of the sharp variation in the computed cross-country correlation beyond the role of trade intensity and the level of financial development. Furthermore, while the latter variables reflect endogenous macroeconomic developments, liberalization episodes are variables directly related to policy decisions undertaken in these countries. Their impact reflects how governments might influence financial interdependence and hence international portfolio selection strategies. In this respect, our study is obviously related to previous papers quoted before such as Bekaert and Harvey (2000). It nevertheless relies on a sound econometric investigation integrating new statistical developments and makes use of alternative measures of liberalization for a large set of emerging countries.

The empirical investigation is conducted on pairs of 25 countries [2] observed at an annual frequency over the 1990-2004 period. Our specification is based on a standard panel data gravity model explaining the evolution of the cross-country correlations. Our results suggest that trade

focus on contagion or interdependence effects among geographical groups as among the Asian (Masih and Masih, 1999) or Latin American (Choudhry, 1997) countries.

[2] An appendix lists the countries, which have been included in the study.

and financial liberalization policies tend to raise the correlation between national stock markets. In turn, since one of the used measure of financial liberalization has been found to be related with international capital flows, this suggests that the home bias observed in international equity portfolios might decrease as developing countries become more integrated.

The paper is organized as follows. Section 2 details the computation of our measure of stock market synchronization. Section 3 presents our methodology and section 4 is devoted to the data issues. Section 5 presents the results and section 6 briefly concludes.

# 2 Methodology

## 2.1 Measuring stock Market Linkages

The study of the determinants of stock market linkages raises the issue of how to measure synchronization. Here, we follow the usual approach adopted in most studies (Pretorius, 2005; Dellas and Hess, 2005; Walti, 2005) by capturing synchronization by the value of the pairwise correlation of stock returns across countries over a given period. Here, we study the evolution of the bivariate correlation at the annual frequency in order to capture the influence of macroeconomic variables (such as trade intensity of growth differential) or policy events (such as trade and financial liberalization).

An important point in measuring correlation is to account for its time-varying feature. It is well known that for a given pair of countries, interdependence obviously varies over time. For instance, correlation across markets is found to be higher during phases of financial crisis and failure to account for that may lead to misleading interpretations (Forbes and Rigobon, 2004). Using average correlations over a particular period might be suited to identify factors such as distance or language similarity that influence only the cross-sectional differences in stock returns but do not vary over time. It prevents however the sound investigation of the role of factors that vary across countries but also over time. In this respect, the study of the impact of reforms that aims at liberalizing trade flows or financial investments requires the use of a time-varying measure of cross-country correlation.

For the sake of computing such a measure, we rely in this paper on the concept of realized moments or more precisely realized correlation. The approach of realized moments has been developed in the area of financial econometrics by several authors (Andersen et al., 2003) as an alternative approach to the use of parametric models such as the GARCH of the Multivariate GARCH models. The general idea is to make use of data at a higher frequency in order to build a consistent estimate of the moments of the distribution of financial returns. In our case, we use daily data on stock returns of the various national stock markets in order to compute annual estimate of the cross-country correlation.

Let us define $p^i_{t,d}$ as the value of the stock index of country $i$ at year $t$ ($t$=1, ..., 15) and day $d$ ($d = 1, ..., D_t$) where $D_t$ is the total number of business days in year $t$. A similar notation holds for country $j$ index, $p^j_{t,d}$. The data are closing quotations for the two markets. From these stock indexes, one can then define daily returns as $r^i_{t,d} = ln(p^i_{t,d}/p^i_{t,d-1}) * 100$. As proposed by Andersen et al. (2003), it is possible to build consistent estimates of annual index volatility using the sum of the squared returns observed during the period :

$$\sigma_{t,i}^2 = \sum_{d=1}^{D_t} \left[ r_{t,d}^i \right]^2 \tag{1}$$

A similar measure applies to the stock index relative to country $i$. [3] Basically, this measure- called realized volatility- captures volatility at a annual frequency using the daily squared returns during that particular year (this is why it is also called integrated volatility). Of course, the realized measure does not restrict to the centered second moment. Using the same approach, one can build a measure of realized covariance between the annual stock returns of country $i$ and country $j$:

$$\sigma_t^{ij} = \sum_{d=1}^{D_t} \left[ r_{t,d}^i * r_{t,d}^j \right] \tag{2}$$

Finally, as a measure of time-varying co-movement between the national annual stock returns, one can define the concept of realized correlation $\rho_y^{ij}$ from the realized volatilities and the realized covariance as (Andersen et al. (2001)) :

$$\rho_{ij,t} = \frac{\sigma_t^{ij}}{\sigma_t^i * \sigma_t^j} \tag{3}$$

It might be desirable to compare the estimates of $\rho_t^{ij}$ through the approach of realized moments with the more traditional measures of comovements. To this aim, Figure 1 plots for a couple of countries pairs the dynamics of our estimates with the correlation coefficients computed over periods of one year. From the visual inspection of the eight panels in Figure 1, two comments are in order. First, in general, both estimates capture the long-run movements in the dependence measures. A rise in the estimate with realized moments is also reflected in the rolling correlation coefficients. This suggests that on the whole, the results of the impact of trade and financial liberalization should not differ too much between the two methods. Such a robustness analysis will be proposed in section 4.2.3. Second, in general, the rolling coefficients of correlation tend to be smoother than the estimate through the method of realized moments. This suggests that the realized moments estimates could capture responses of the stock market comovements to liberalization that are not captured by the traditional coeeficients.

The pairwise realized correlations are estimated for the $\frac{25\times24}{2}$ pairs of countries and for the 15 years of data. This leads to a balanced panel data set of 4500 correlations that will be investigated in the next section. For easiness of exposition, the value of the pairwise correlation between $i$ and $j$ at time (year) $t$ is noted as $\rho_{ij,t}$. Since correlations are bounded between -1 and 1, it might be interesting for estimation purposes to reexpress the estimates on a continuous sale ranging from $-\infty$ to $\infty$. Therefore we use a Fisher-Z trasformation of $\rho_y^{ij}$ :

$$\overline{\rho}_{ij,t} = ln(\frac{1 + \rho_{ij,t}}{1 - \rho_{ij,t}}) \tag{4}$$

[3] Notice that in equation (1), we assume that $E(r_{t,d}^i r_{t,d-1}^i) = 0$. This is an usual assumption for high frequency financial data. See Andersen et al.(1999) on this point.

## 2.2 Econometric Issues

Basically, the empirical approaches used to gauge the degree of financial integration might be divided in two different strands.

A first type of approach look at the strength of the relationship between the dynamics of the national stock price indexes and some measure of the common shocks. A first subset of empirical studies relies on latent factor models to measure the extent of financial integration (Emiris, 2006 or Bruneau and Flageolet, 2005 for instance). In this approach, the share of the variation of national stock indices due to the most important common factor(s) across countries of interest provides a direct measure of the degree of integration of this particular country. If the model is made dynamic, one can also provide the timings of increased integration. An example of a second subset of approaches is provided by Fratzscher (2001) using a multivariate GARCH modelling. In this framework, the degree of integration is measured by the amplitude of the response of the national stock returns to European and US (daily) stock market innovations. These approaches might be useful when it is cumbersome to build explicit measures of integration or when there is a strong presumption that the legal constraints to free movements might be easily circumvented.

As an alternative, when relevant measures of trade and financial constraints are available, one can estimate explicitly the relationship between some measure of co-movements and these constraints. The present paper follows this approach. More precisely, our econometric approach is based on the empirical estimation of a gravity type of model allowing for the combination of the cross-sectional and time series information. While the gravity model has been extensively in the empirical trade literature, it has been found to offer a good starting point in explaining the evolution of cross-market correlations (Walti, 2005; Flavin et al., 2002). However, this model has often been estimated using pure cross section data (Dellas and Hess, 2005) or by pooling the data (Flavin et al., 2002; Walti, 2005), leaving aside the time dimension of the integration process. In contrast to these approaches, we explicitly account for the issue of unobserved heterogeneity by estimating a panel model with cross-section and period specific effects:

$$\overline{\rho}_{ij,t} = \alpha + \lambda_{i,t} + \lambda_{j,t} + X'_{ijt}\beta + \delta_{ij} + \gamma_t + \epsilon_{ijt}, \qquad (i \times j) \in (1,..,n)^2, \qquad i < j \qquad t = 1, ..., T, \quad (5)$$

where $\overline{\rho}_{ij}$ is the realized correlation between stock markets in countries $i$ and $j$ at time $t$, and $X_{ijt}$ is a vector of exogenous regressors, and $\epsilon_{ijt}$ are the error terms for cross-sectional units observed for dated periods. The $\alpha$ parameter represents the overall constant in the model, the $\lambda_{i,t}$ and$\lambda_{j,t}$ are variables (such as the exchange regime classification) that are country-specific and varies over time, while the $\delta_{ij}$ and $\gamma_t$ represent cross-section or period specific effects, which can be random or fixed. Our setup is also characterized by a cross-section dimension ($n$) larger than the time dimension ($T$). Such dimensions are quite unusual for a macroeconomic setup, for which small cross-section dimensions are often the norm.

The introduction of cross-section and period specific effects is important in several respects. First, failure to control for unobserved heterogeneity might result in biased estimators and in turn can lead to misleading conclusions. Second, the inclusion of cross-section specific effects ($\delta_{ij}$) will capture in a parsimonious way the influence of non time-varying factors. A couple of factors of this type have been found to influence cross-market correlations, such as geographical distances, common border or synchronicity in trading hours (Flavin et al., 2002). To a certain extent, the cross-section effects will also capture the role of time-varying but highly persistent variables such as similarity in industrial structure which has also been included in previous studies (Dellas and Hess, 2005; Walti, 2005). Likewise, the time-specific effects ($\gamma_t$) allows us to capture the influence of common shocks hitting all countries. The need of controlling for common shocks has been stressed by several authors. Chinn and Forbes (2004) for instance include global variables such as world interest rates. Our approach might be seen as an alternative to the explicit choice of global variables. Cross-market correlations are known to increase in times of financial crisis with large contagion effects.

In our study, the inclusion of time dummies might also be important since our investigation period includes the occurrence of large financial crisis such as the Asian or the Russian crises. To sum up, the inclusion of both effects allows to focus on the role of specific $X_{ijt}$ variables, such as financial integration and trade liberalization reforms, without failing to control for structural factors that determine the size of the cross-market correlations.Another global unobservable variable that could be captured by the time dummies is technology. Technological advances have led to an increase of cross-border financial flows, which in turn can induce an overall increase in cross-country stock market comovement. Once again, assuming that the effects of these omitted variables is identical across countries, the introduction of these time dummies significantly reduces the scope of mispecification in the regression models. To sum up, the inclusion of both effects allows to focus on the role of specific $X_{ijt}$ variables, such as financial integration and trade liberalization reforms, without failing to control for structural factors that determine the size of the cross-market correlations.

Before going further with the estimation of equation (5), preliminary Haussman tests have to be performed in order to determine wether the cross-section or/and the period specific effects are random or fixed. The results indicates that both effects appear to be fixed.

With respect to the estimation *stricto sensu*, this model can be estimated via GLS. Nevertheless, as we suspect that cross-section dependence could be present [4], the feasible GLS approach is applied. It also has the advantage of tackling potential heteroscedasticity. This is important because accounting for time-series heteroscedasticity has been stressed by Forbes and Rigobon (2002) for using correlation as a measure of c-movement. The use of Newey-West standard errors allows to minimize the influence of financial crisis on our estimation results.

To account for potential dynamics, model (5) is augmented to integrate an autoregressive component and hence becomes:

$$\overline{\rho}_{ij,t} = \alpha+\lambda_{i,t}+\lambda_{j,t}+\overline{\rho}'_{ij,t-1}\beta_1+X'_{ij,t}\beta_2+\delta_{ij}+\gamma_t+\epsilon_{ijt}, \qquad (i\times j) \in (1,..,n)^2, \qquad i < j \qquad t = 1,...,T. \tag{6}$$

From this specification, it is possible to distinguish the short-run ($\alpha$, $\beta_2$) and the long-run coefficients: $\frac{1}{1-\beta_1}.(\alpha, \beta_2)$. As we reject the presence of a unit root in the realized correlation variable, $\beta_1 < 1$. Several studies have shown that the estimation of the autoregressive parameter could be biased (see Phillips and Sul, 2002), even though the presence of an exogenous variable could reduce this bias.[5] Several alternative methods are then proposed: Phillips and Sul (2003) derives formulae to correct for the bias in dynamic panel estimation with fixed effects and cross-

[4] For each specification, cross-section dependence is tested via Pesaran's test (2004).

[5] Nickell (1981) shows that the bias is increasing in the size of the ratio of $n$ over $T$. Given the large number of cross-sections (300) compared to the number of periods (15), the potential bias should be rather limited here.

section dependence. Nevertheless, in the case of cross-section dependence, the correction is not easy because of the presence of random elements. The second approach consists in specifying the cross-section dependence, via a spatial weight matrix (see Anselin, 2000 or Hahn and Kuersteiner, 2002). This method is popular because the spatial matrix is exogenous, inherited from geographical information common to anybody. Nevertheless, this approach is feasible only if $T$ tends to infinity faster that $n$, which is obviously not our case. The third method has been proposed by Arrelano and Bond (1991) and consists of GMM estimator. The instrument space is composed of the past of $Y_{it}$ and $X_{it}$. Nevertheless, this estimator may be biased in the case of cross sectional dependence (see Phillips and Sul, 2003, and Hahn and Kuersteiner, 2002). Therefore, we decide to use the FGLS estimator [6], which controls for cross-section dependence and heteroscedasticity. It is obvious that even if the same estimator is used in the case of the static and the dynamic cases, they will differ with respect to the weights used for the estimation.

# 3 Data Issues

Since our main interest lies in the impact of trade and financial liberalization, we focus on the way these developments are measured in our empirical analysis.

## 3.1 Capturing financial liberalization

As analysed by Bekaert and Harvey (2003), there are several ways to capture financial liberalization episodes. One can use the dates of announcements made by the governments themselves. A major drawback of using these dates is that in practice administrative constraints and capital controls might still be binding even after the official announcements, so that no real openness to foreign investments takes place. Furthermore, since the behaviour of international investors depends on the quality of the signal sent by the financial authorities, this raises the issue of the credibility of such announcements.[7] Tackling this specific point by estimating the breaks from the data generating process of a couple of financial time series, Bekaert et al. (2003) find that the endogeneous dates of financial liberalization are usually later than official dates. As an alternative, one can use the timing at which investment funds were launched in the country. This might reflect in a more realistic way the effective context regarding financial liberalization. In this respect, we use the dates of provided by Bekaert and Harvey (2002) and build bilateral liberalization indices.[8] More precisely, we build a dummy variable $f_{ij,t}$ taking 1 if countries $i$ and $j$ were both considered liberalized at time $t$ and 0 otherwise.

## 3.2 Capturing trade liberalization

The same procedure can be used to build bilateral trade liberalization indices. Here we use the trade liberalization dates provided by Sachs and Warner (1995) which are based on a set of

[6]This estimator has been used in the case of dynamic panels with cross section dependence in several studies, i.e. Peersman and Smeets (2005).

[7]A good example is provided by the case of Morocco. On the basis of the official liberalization measure, Morocco has always been liberalized since 1990. In spite of that, there has not been any domestic investment fund launched so far. This raises the question of the remaining impediments to the launch of such a fund.

[8]The orrelation between both measures of financial liberalization amounts to about 0.15. This reflects that the choice of the liberalization date is of primary importance and that there is a significant discrepancy between the effective data and the official ones.

criteria such as average tariffs or quotas. In contrast to trade intensity, which tends to reflect the importance of trade relationships, the trade liberalization dummies will capture more the impact of policy reforms in that area. Beyond their impact on actual trading intensity, these policy reforms might also exert a signalling impact on the behaviour of international investors. Compared to "closed" economies, asymmetric information between two trading partners is reduced. In turn, the expected decrease in asymmetric information might induce investors to diversify their portfolio internationally.

### 3.3 Trade intensity

Following Pretorius (2002), we build a relative trade intensity measure reflecting the importance for a given country of the trading partners of a given country. More precisely, we use the sum of the shares of the trading partner in the total trade of each country:

$$T_{ij,t} = (\frac{Z_{ij,t} + M_{ij,t}}{Z_{i,t} + M_{i,t}}) + (\frac{Z_{ij,t} + M_{ij,t}}{Z_{j,t} + M_{j,t}}) \tag{7}$$

where $Z_{ij,t}$ and $M_{ij,t}$ are the value of exports and imports from country $i$ to country $j$ while $Z_{i,t}$ and $M_{i,t}$ are the value of total exports and imports of country $i$. Trade intensity tends to increase for a large majority of country pairs over our sample period (1990-2004). This reflects the tendency towards higher globalization. As a result, the trade intensity measures are highly correlated with a time trend. For the purpose of panel estimation, this is an important point since this induces a high degree of collinearity in models with fixed time-specific effects. Therefore, to reduce the amount of correlation, we use in these models the change in $T_{ij,t}$ rather than the level. In contrast, in models with particular time dummies capturing specific common shocks such as the major financial crisis, the level of trade intensity is used in the regressions.

### 3.4 Correlation between liberalization measures

For the sake of interpretation of the results, it is interesting to look at the relationships between liberalization measures. As for the correlation between trade liberalization and trade intensity, one can expect some positive link between both concepts. Nevertheless, the correlation of trade intensity with the Sachs and Warner measure indeed amounts to 0.01, reflecting a very loose relationship between both variables. Several explanations are in order here. First, there is obviously a delay between liberalization and the increase in trade flows between countries. Second, most trade agreements seem to be implemented when the potential gains are the biggest. In other terms, there is less need to make explicit bilateral agreements when trade intensity is already high.

The data also reveals that there is a moderate relationship between trade liberalization and financial liberalization in our sample. The correlation of trade liberalization with actual (resp. official) dates of financial liberalization amounts to 0.16 (resp. 0.13). This might reflect that the willingness to liberalize trade goes moderately hand in hand with the willingness to open the capital account. This means that for a subset of countries, the economic policy in terms of liberalization applied equally to the real and the financial side of the economy. Nevertheless, for a remaining number of countries, the political decisions were taken independently or at least involve significant delays. In this respect, the introduction of a time dimension in the sample is important here.

The observed weak relationship between the explanatory variables of main interest has several implications. First, from an econometric point of view, the scope of collinearity in our regression models is quite limited, which in turn increase the quality of the estimates. Second, the data reveals that both concept of trade relationships need to be taken into account to capture the impact of trade liberalization on the cross-country correlation in a consistent way. In particular, it seems that both measures do not reflect some redundant information concerning bilateral trade relationships.

### 3.5 Exchange rate regime

Exchange rate commitments are classified by the IMF in the international monetary fund's annual report on exchange rate arrangments and exchange restrictions and label as the *de jure classification*. Nevertheless, several authors (in particular Levi-Yeyati and Sturzenegger, 2003) has shown that some countries do not respect their commitments and have justified the use of *'de facto classifications*. Following Levi-Yeyati and Sturzenegger 2003, a 3-regime classification [9] is considered. [10] Unfortunately, there are, to the best of our knowledge, no available data measuring the extent of variability of the bilateral exchange rate. [11] This leads to 3 possible values for a $X_{ij,t}$ type of variable ($X_{ij,t} = 1, 2$ or 3) capturing the degree of variability of the exchange rate between country $i$ and country $j$ at time $t$. More precisely, the variable takes the value of the most flexible system of the two countries. As an example, if country $i$ is considered in a flexible exchange rate regime while country $j$ is in a fixed exchange rate regime, the variable takes the value of the flexible exchange rate system, i.e. $X_{ij,t} = 1$. [12]

### 3.6 Other variables

As additional control variables, we introduce macroeconomic determinants that are assumed to influence the evolution of stock prices. Following previous empirical studies (Pretorius, 2002, for instance), we include interest rate differentials. The interest rates are 3 months interest rate and we use average values over the year. We also include the inflation differential computed as the difference in the growth rate of consumption price indices. Last but not least, we also introduce the real GDP growth differential.

The panel data regression model allows time dummies to be introduced. These time dummies can easily capture the impact of major shocks to correlations such as major financial crisis. We proceed with two separate investigations. In a first model, we include time dummies for the years 1995, 1997 and 1998, which aim at capturing the impact of the Mexican crisis, the Asian crisis and the Russian crisis, respectively. In the second model, we estimate a model in which all time dummies are included, leading to a two-way error component model. We also include bilateral geographical area(s) dummies taking 1 if both countries belong to the same sub-continent. More

[9] Floating, intermediate and fix exchange rate

[10] We also consider a variable representing the 5-regime classification, but it turns out to be insignificant in all experiments. Results are available upon request from the authors.

[11] Of course, it is always possible to build measures of variability based on the historical exchange rate fluctuations. Nevertheless, mapping these measures into bilateral regime classification would require the use of arbitrary criteria. As an alternative, we prefer to use regime classifications based on previous classifications.

[12] While being less intuitive, we also consider the variable taking the value of the less flexible system of the two countries. Results are not modified by this variable which turns out to be not significant.

precisely, we build that only for Asian and Latin American countries since most of the emerging countries come from these areas.

# 4 Results

## 4.1 Benchmark results

Tables 1 and 2 report the estimation results of our panel gravity model. In contrast to the estimated models of Table 1, the specifications of Table 2 allows for some dynamics in the co-movement through the inclusion of an $AR(1)$ component.[13] The dynamic specification is supported by the significant AR1 term but also by the fact that the absence of autocorrelation is not rejected by traditional Breusch-Pagan and z-tests. However, our main results regarding the impact of trade and financial liberalization are qualitatively identical regardless the used specification.

Results of the specification tests reported at the bottom of the tables indicate the nature of the estimated processes and suggest the use of particular estimation techniques. First, Hausman tests tend to support the use of fixed individual and time effects rather than random ones. All time dummies are significant at usual significance levels, suggesting that accounting for unobserved time heterogeneity is important. It indicated that all the "fixed" variables, such as geographical distance or industrial similarities which are generally found significant in cross-section gravity models, are included in this fixed effect term. With both fixed effects included in the model, we find a significant AR1 component around 0.09.[14] Such an estimate suggests that it takes on average 11 years to reach the steady state equilibrium. Estimating the models with OLS by pooling the data, we find an $AR(1)$ component close to 0.04. It indicates that the dynamics might be quite different when failing to account for unobserved heterogeneity: OLS underestimate the adjustment time to long-run equilibrium. Finally, the Pesaran test suggests the presence of cross-sectional dependence. This justifies the use of FGLS estimators even with a dynamic specification. Besides, we notice that the explanatory power of our panel model (static or dynamic) is much higher than that found in traditional cross-section gravity models.

Estimation results suggest that, in general, macroeconomic variables such as growth and the inflation differential are poorly related to stock return co-movements. This results in line with the findings of Canova and De Nicolo (1997). They find that the European stock returns are explained by US inflation and real variables rather than by domestic variables. Foreign variables can indeed drive the correlations as they are good predictors of future domestic activity. In our analysis, since US variables are common to all countries, their effect will be captured through the time dummies. In the full model, inflation differentials tend to be positively related to stock co-movements, which is rather counter-intuitive. Our measures of the exchange rates regimes do not provide any significant explanatory power to the co-movement of stock prices. This might be due partly to the fact that the classification regime variables are country-specific and not pair-specific. In contrast, the time dummies capturing the occurrence of the major crises tend to be highly related to correlation. It is found that the Asian and Russian crises led to a general increase in stock market co-movement, which is consistent with some well-known stylized facts (Forbes and

[13] Autoregressive components of higher orders turn out to be insignificant.

[14] This value turns out to be robust to the specification of the model and to the use of the Fisher-Z transformation or not.

Table 1: Impact on stock market correlations: static approach

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| *Constant* | 0.056 | 0.073*** | 0.080*** | 0.161*** | 0.167*** |
| | (0.040) | (0.027) | (0.008) | (0.057) | (0.027) |
| *financial liberalization* | 0.073 | 0.093** | 0.092*** | 0.206*** | 0.201*** |
| | (0.052) | (0.037) | (0.020) | (0.026) | (0.035) |
| *trade liberalization* | 0.036*** | 0.026** | 0.017 | 0.051*** | 0.035** |
| | (0.013) | (0.012) | (0.017) | (0.015) | (0.018) |
| *trade intensity* | 0.078* | 0.079*** | 0.150* | 0.154*** | 0.328** |
| | (0.046) | (0.034) | (0.091) | (0.059) | (0.183) |
| *Latin America* | 0.016 | 0.013* | 0.009 | 0.028** | 0.019** |
| | (0.010) | (0.008) | (0.008) | (0.010) | (0.011) |
| *Asia* | 0.008 | - | - | - | - |
| | (0.006) | | | | |
| *dummy 95* | -0.059*** | -0.059*** | - | -0.126*** | - |
| | (0.011) | (0.013) | | (0.018) | |
| *dummy 97* | 0.094*** | 0.093*** | - | 0.199*** | - |
| | (0.011) | (0.012) | | (0.018) | |
| *dummy 98* | 0.088*** | 0.088*** | - | 0.189*** | - |
| | (0.011) | (0.012) | | (0.022) | |
| *differential int rates* | 0.056 | - | - | - | - |
| | (0.242) | | | | |
| *differential inflation* | 0.012 | - | - | - | - |
| | (0.010) | | | | |
| *differential growth* | 0.001 | - | - | - | - |
| | (0.001) | | | | |
| *exchange rate regime* | -0.006 | - | - | - | - |
| | (0.012) | | | | |
| Number obs | 4108 | 4500 | 4200 | 4500 | 4200 |
| $R^2$ | 0.398 | 0.381 | 0.423 | 0.365 | 0.423 |
| LM B-P test | $< 1\%$ | $< 1\%$ | $< 1\%$ | $< 1\%$ | $< 1\%$ |
| z-test | $< 1\%$ | $< 1\%$ | $< 1\%$ | $< 1\%$ | $< 1\%$ |
| Hausman Test | $< 1\%$ | $< 1\%$ | $< 1\%$ | $< 1\%$ | $< 1\%$ |
| Pesaran test | $< 1\%$ | $< 1\%$ | $< 1\%$ | $< 1\%$ | $< 1\%$ |

Estimated model:$\rho_{ij,t} = \alpha_{ij} + \phi' x_{ij,t} + \gamma' z_t + \epsilon_{ij,t}$. White standard errors of estimates are in parentheses. All models include fixed individual effects. Individual effects not reported to save space.***, ** and * denote the significance of estimates at 10, 5 and 1 percent levels respectively.

Figures associated to the Hausman homogeneity test report the p-value of the test with null hypothesis of random effect against the alternative of fixed effect in the cross and the time dimension. Similar results (not reported here) are obtained when disentangling both dimensions. Figures associated to the Pesaran test report the p-value of the Pesaran (2004) test with null hypothesis of no cross sectional dependence.

Figures associated to the $LMB-P$ (Breusch-Pagan) reports the p-value of the test with the null hypothesis of no $1^{st}$ order serial correlation.

Figures relative to the $z-test$ (Fisher) reports the p-value of the test with the null hypothesis of no $1^{st}$ order serial correlation.

Figures associated to the $1^{st}$ order autocorrelation report the p-value of the test with null hypothesis of no vs AR(1) errors.

Columns (3) and (5) include fixed time effects (not reported here to save space) and capture trade intensity by change in bilateral trade rather the by the level.

Results in columns (1) to (3) are estimated using the unadjusted value of correlation $\rho_{ij,t}$, while in columns (4) and (5), the adjusted values $\overline{\rho}_{ij,t}$ are used.

Rigobon, 2004). It is also found that Latin American markets tend to be, in general, more related. This is consistent with the fact that common borders, and to a lesser extent common language(s) between countries tend to lead to higher co-movement in stock prices (Flavin et. al, 2002). Finally, it turns out that exchange rate regime is not related to stock exchange co-movement. This result contradicts the conclusion of Walti (2005) for developed countries.

Trade intensity has been considered one of the main channels through which contagion effects of financial crises take place (Glick and Rose, 1999; Chinn and Forbes, 2004). Nevertheless, the effect of trade intensity on stock return co-movement remains ambiguous in empirical studies, as shown by the results of Dellas and Hess (2005) or those of Pretorius (2002). We find moderate support for a positive effect of trade intensity on stock return co-movement. In the full dynamic specification (model (2)), trade intensity is significant only in the parsimonious specifications, suggesting a moderate role for trade linkages. Nevertheless, trade liberalization reforms are found to increase stock market co-movements in a less ambiguous way. Further investigation shows that the trade liberalization dummies are weakly (and positively) correlated with trade intensity : the unconditional correlation coefficient amounts to 0.013. This might be a reflection of the fact that trade liberalization reforms often take time to exert significant effects on the size of trade. The fact that such reforms exert a significant impact beyond the traditional role of trade intensity suggests the existence of a signalling effects for international investors. Reforms aimed at increasing the extent of trade relationships might lead to a reduction in the degree of asymmetric information, favoring in turn the size of cross-border investment.

Financial liberalization is also found to have a positive and significant impact on stock market correlation. This impact is highly significant for all specifications, a result which holds regardless of the measure of the stock market co-movement. This result is in line with those of Walti (2005). Using the $\rho_{ij,t}$ as the dependant variable (rather than $\overline{\rho}_{ij,t}$ whose results are in Tables 1 and 2), we find that bilateral correlations between liberalized countries are higher on average by 0.07 to 0.10 compared to countries in which international investment is constrained. Since our financial liberalization variable captures dates of launch of at least one country investment fund, this impact might be driven by both a signalling effect but also by a direct increase in bilateral investment flows.

Our results might have also some implications with respect to the home country bias puzzle. Of course, further investigation is required to refine these implications, for instance by using data on cross-border investment flows (which are difficult to obtain, especially for emerging countries). Under the assumption that the liberalization reforms tend to lead to an increase of these bilateral investment flows (which in turn would lead to a higher synchronization of asset prices across countries), this might suggest that financial and trade protections explain part of this puzzle. Likewise, under the same assumption, the fact that trade intensity ultimately leads to an increase in the correlation, this is consistent with the fact asymmetric information is also an important factor of the home country bias. Increase in trade relationships can decrease the amount of asymmetric information, inducing investors to buy and sale foreign stocks.

## 4.2 Robustness Analysis

In this section, we proceed to robustness checks in three different directions. First, we depart from the use of dummies to capture dates of financial openness and consider recent alternative

Table 2: Impact on stock market correlations: dynamic approach

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| *Constant* | 0.003 | 0.029*** | 0.070*** | 0.064*** | 0.142*** |
| | (0.013) | (0.001) | (0.012) | (0.023) | (0.027) |
| *lagged correlation* | 0.073*** | 0.082*** | 0.097*** | 0.086*** | 0.119*** |
| | (0.016) | (0.015) | (0.016) | (0.015) | (0.018) |
| *financial liberalization* | 0.138*** | 0.127*** | 0.089*** | 0.271*** | 0.192*** |
| | (0.016) | (0.014) | (0.016) | (0.031) | (0.035) |
| *trade liberalization* | 0.041*** | 0.029*** | 0.0149** | 0.058*** | 0.033** |
| | (0.008) | (0.008) | (0.009) | (0.017) | (0.018) |
| *trade intensity* | 0.073** | 0.079*** | 0.007 | 0.151** | 0.327*** |
| | (0.042) | (0.030) | (0.005) | (0.062) | (0.181) |
| *Latin America* | 0.012** | 0.008 | 0.009* | 0.016 | 0.015 |
| | (0.006) | (0.005) | (0.005) | (0.010) | (0.011) |
| *Asia* | 0.007 | - | - | - | - |
| | (0.006) | | | | |
| *dummy 95* | -0.053*** | -0.055*** | - | -0.116*** | - |
| | (0.009) | (0.009) | | (0.018) | |
| *dummy 97* | 0.100*** | 0.097*** | - | 0.207*** | - |
| | (0.008) | (0.009) | | (0.018) | |
| *dummy 98* | 0.085*** | 0.081*** | - | 0.174*** | - |
| | (0.010) | (0.009) | | (0.021) | |
| *differential int rates* | 0.039 | - | - | - | - |
| | (0.029)* | | | | |
| *differential inflation* | 0.194** | - | - | - | - |
| | (0.088) | | | | |
| *differential growth* | 0.000 | - | - | - | - |
| | (0.001) | | | | |
| *Exchange rate regime* | -0.004 | - | - | - | - |
| | (0.007) | | | | |
| Number obs | 3878 | 4200 | 4200 | 4200 | 4200 |
| $R^2$ | 0.407 | 0.404 | 0.429 | 0.391 | 0.439 |
| LM B-P test | 61.06% | 15.57% | 56.62% | 4.28% | 67.03% |
| z-test | 60.27% | 17.39% | 56.14% | 5.54% | 65.86% |
| Hausman Test | $< 1\%$ | $< 1\%$ | $< 1\%$ | $< 1\%$ | $< 1\%$ |

Estimated model:$\rho_{ij,t} = \alpha_{ij} + \delta\rho_{ij,t} + \phi' x_{ij,t} + \gamma' z_t + \epsilon_{ij,t}$. White standard errors of estimates are in parentheses. All models include fixed individual effects. Individual effects not reported to save space.***, ** and * denote the significance of estimates at 10, 5 and 1 percent levels respectively.
Figures associated to the Hausman homogeneity test report the p-value of the test with null hypothesis of random effect against the alternative of fixed effect in the cross and the time dimension. Similar results (not reported here) are obtained when disentangling both dimensions.
Figures associated to the $LMB-P$ (Breusch-Pagan) reports the p-value of the test with the null hypothesis of no $1^{st}$ order serial correlation.
Figures relative to the $z-test$ (Fisher) reports the p-value of the test with the null hypothesis of no $1^{st}$ order serial correlation.
Figures associated to the $1^{st}$ order autocorrelation report the p-value of the test with null hypothesis of no vs AR(1) errors.
Columns (3) and (5) include fixed time effects (not reported here to save space) and capture trade intensity by change in bilateral trade rather the by the level.
Results in columns (1) to (3) are estimated using the unadjusted value of correlation $\rho_{ij,t}$, while in columns (4) and (5), the adjusted values $\overline{\rho}_{ij,t}$ are used.

measures. Second, we estimate the dynamic panel model with the well known GMM estimators developed by Arrelano and Bond (1991) rather than relying on the FGLS estimator. Finally, we use traditional rolling correlations as a measure of stock market co-movement instead of the realized correlation.

#### 4.2.1 Alternative measures of financial liberalization

Up to now, we have used a particular measure of financial integration based on dates of capital liberalization. As claimed by Bekaert and Harvey (2002) these dates correspond to effective opening to foreign capital flows and should capture the major developments in terms of financial liberalization. Nevertheless, as claimed by Edison and Warnock (2003), the use of dummies exhibits some drawbacks. The use of dummies implicitly assumes that financial liberalization takes place as one-shot event. Once liberalized, capital flows are completely free to move in and out of the country. This assumption is not in line with what is observed in some countries such as Malaysia for which capital controls have been restored after a long period of liberalization. Furthermore, there is some evidence that financial liberalization is a gradual process which cannot be captured by dummies such as the ones proposed by Bekaert and Harvey (2002).

To overcome the limitation of our first measure of financial liberalization, we build bilateral measures based on the measure proposed by Edison and Warnock (2003). [15] The Edison and Warnock data provides a quantitative measure of the proportion of each country's equities held by foreigners. We follow Edison and Warnock (2003) by taking one minus this proportion. This variable might be used as a measure of capital controls and provides an assessment of the degree of financial liberalization. The main drawback of using this measure in our analysis is that these measures exist only over the 1990-2000 period and are not available for 6 countries included in our initial sample. [16] Therefore, we will use these results mainly as a robustness check to assess the reliability of the results obtained with our large set of countries and with the the dummy measures of financial liberalization.[17]

In order to compute bilateral measures of financial liberalization, we consider two alternative measures. First, we take the minimum of the original data. This assumes that the capital and investment flows between the two countries are subject to the constraints of the most restrictive country at that time. This measure is labeled as Edison-Warnock 1 in the subsequent tables of results. As an alternative, we also compute the product of the Edison-Warnock measures. This is labeled as Edison-Warnock 2 hereafter. It should be stressed that negative coefficients are consistent with the previous positive estimates of the financial liberalization of Tables 1 and 2 in the sense that it is expected that less capital control tend to increase the degree of stock market co-movements. Given the availability of the original measures, we end up with a sample of less than 3,000 observations instead of 4,500 in the benchmark regressions. Table 3 report the results

[15] Several other measures for financial liberalization are available see Miniane, 2004, Kaminsky and Schmukler (2003). Nevertheless, using such measures would lead to reduce our sample by more than 10 countries. Robustness check would then become unappropriate.

[16] These countries are Columbia, Israel, Japan, Kenya, Mauritius and Singapore. The Edison and Warnock database are only available at a monthly frequency and they have been transformed to an annual frequency via yearly averages.

[17] Looking at the correlation between the Edison-Warnock measure and the liberalization dates used in the benchmark regressions shows that the relationship is quite weak for some countries. As a result, pooling the data, the global correlation between the alternative measures is close to zero, which emphasizes the need for a robustness analysis.

for the dynamic specifications (complete and parsimonious specifications), with the three time dummies capturing the occurrence of financial crises.[18]

Results reported in Tables 3 show that the use of an alternative measure of financial liberalization leads to the same results. In particular, we find that the openness of financial markets to foreign investors tends to increase the co-movement of the domestic stock market with its foreign counterpart. Interestingly, the results obtained for the impact of trade liberalization are also very in line with the previous ones.

On the whole, the results obtained in tables 3 tend to confirm the previous findings. In general, financial liberalization is found to lead to an increase in the stock prices co-movement. A small exception is found when using time dummies for each period to capture the impact of common shocks. While negative as before, the coefficients are not significantly different from zero. This results is due to the strong correlation between the Edison-Warnock data and the time trend, suggesting that in general, there was a gradual liberalization process in the countries included in the sample.[19] The high significance of the financial liberalization obtained when a subset of these time dummies only are included reflect that this result is driven mainly by the joint statistical features of both variables. As for the impact of trade liberalization and trade intensity, the results of tables 3 and 4 are very in line with those obtained in the benchmark regressions. Therefore, one can reasonably conclude that our previous results are robust to some alternative measure of financial liberalization.

As shown by Edison and Warnock (2003), the decrease in the degree of capital controls went hand in hand with strong increases in capital inflows to those countries. Although these flows are not included in our analysis (the data are in general not available for most of the countries in the sample) this finding suggests that the increase in the correlation might be driven by international financial movements between the countries. This is consistent with the idea that part of the home bias puzzle is related to the existence of international barriers of capital movements. The inclusion of this missing variable and the estimation of a full system rather than a reduced form equation is a promising avenue of future research but is clearly beyond the scope of this paper.

#### 4.2.2 GMM estimation

As argued in section 3, the use of FGLS estimators to estimate equation (6) rather than the popular GMM method (Arrelano and Bond, 1991) is relevant in the case of cross-sectional dependence. Since our dependent variable is a measure of stock price comovement, one can expect the degree of cross-sectional dependence to be high. Nevertheless, the use a dynamic model opens the case for some bias of the estimates (the so-called Nickell bias). Although this bias should be limited given the low dimension of $T$, one way to get rid of the Nickell bias is to use GMM. In this section, we reestimate model (6) using the three alternative measures of liberalization to assess the robustness of the results obtained with FGLS. The results are reported in Table 4.

The results of Table 4 show that the positive impact obtained for trade and financial liberalization cannot be ascribed to the use of FGLS rather than GMM estimators. Correcting for the Nickell bias by the specific use of GMM estimators leads to the same qualitative results. This

[18] The results with the static model can be obtained upon request.

[19] The high degree of collinearity is also reflected by the sharp increase in the standard deviation of the estimated parameter.

Table 3: Impact on stock market correlations: dynamic approach

| | Edison-Warnock 1 | | Edison-Warnock 2 | |
|---|---|---|---|---|
| | (1) | (2) | (1) | (2) |
| *Constant* | 0.104*** | 0.116*** | 0.110*** | 0.118*** |
| | (0.010) | (0.009) | (0.011) | (0.019) |
| *lagged correlation* | 0.032** | 0.052*** | 0.031* | 0.024* |
| | (0.016) | (0.017) | (0.014) | (0.018) |
| *financial liberalization* | -0.138*** | -0.132*** | -0.138*** | -0.132*** |
| | (0.022) | (0.021) | (0.026) | (0.024) |
| *trade liberalization* | 0.043*** | 0.033*** | 0.048*** | 0.034*** |
| | (0.008) | (0.003) | (0.008) | (0.008) |
| *trade intensity* | 0.196*** | 0.220*** | 0.199** | 0.222*** |
| | (0.083) | (0.088) | (0.106) | (0.116) |
| *Latin America* | 0.003 | 0.005 | 0.001 | 0.001 |
| | (0.005) | (0.005) | (0.005) | (0.006) |
| *Asia* | 0.012*** | - | 0.011*** | - |
| | (0.002) | | (0.003) | |
| *dummy 95* | -0.037*** | -0.036*** | -0.038** | -0.037*** |
| | (0.012) | (0.012) | (0.012) | (0.012) |
| *dummy 97* | 0.116*** | 0.116*** | 0.115*** | 0.117*** |
| | (0.011) | (0.011) | (0.011) | (0.011) |
| *dummy 98* | 0.123*** | 0.123*** | 0.124*** | 0.123*** |
| | (0.013) | (0.013) | (0.014) | (0.014) |
| *differential int rates* | -0.012 | - | -0.002 | - |
| | (0.035) | | (0.003) | |
| *differential inflation* | -0.002 | - | -0.001 | - |
| | (0.003) | | (0.001) | |
| *differential growth* | -0.000 | - | -0.000 | - |
| | (0.001) | | (0.001) | |
| *Exchange rate regime* | -0.005 | - | -0.004 | - |
| | (0.007) | | (0.006) | |
| Number obs | 2516 | 2719 | 2516 | 2719 |
| $R^2$ | 0.775 | 0.625 | 0.745 | 0.663 |
| LM B-P test | 94.77% | 35.82% | 62.20% | 37.40% |
| z-test | 95.99% | 36.82% | 61.34% | 38.99% |
| Hausman Test | $< 1\%$ | $< 1\%$ | $< 1\%$ | $< 1\%$ |

Estimated model:$\rho_{ij,t} = \alpha_{ij} + \delta\rho_{ij,t} + \phi' x_{ij,t} + \gamma' z_t + \epsilon_{ij,t}$. White standard errors of estimates are in parentheses. All models include fixed individual effects. Individual effects not reported to save space.***, ** and * denote the significance of estimates at 10, 5 and 1 percent levels respectively.
Figures associated to the Hausman homogeneity test report the p-value of the test with null hypothesis of random effect against the alternative of fixed effect in the cross and the time dimension. Similar results (not reported here) are obtained when disentangling both dimensions. Figures associated to the Pesaran test report the p-value of the Pesaran (2004) test with null hypothesis of no cross sectional dependence. Figures associated to the $LMB-P$ (Breusch-Pagan) reports the p-value of the test with the null hypothesis of no $1^{st}$ order serial correlation.
Figures relative to the $z-test$ (Fisher) reports the p-value of the test with the null hypothesis of no $1^{st}$ order serial correlation.
Figures associated to the $1^{st}$ order autocorrelation report the p-value of the test with null hypothesis of no vs AR(1) errors.

Table 4: Impact on stock market correlations: dynamic approach with GMM estimation

| | Baseline model | Edison-Warnock 1 | Edison-Warnell 2 |
|---|---|---|---|
| *lagged correlation* | 0.006 | -0.019* | 0.046*** |
| | (0.013) | (0.010) | (0.011) |
| *financial liberalization* | 0.104*** | -0.250*** | -0.227*** |
| | (0.049) | (0.022) | (0.015) |
| *trade liberalization* | 0.520*** | 0.833*** | 0.833*** |
| | (0.108) | (0.104) | (0.082) |
| *trade intensity* | 0.143*** | 0.105*** | 0.080*** |
| | (0.014) | (0.011) | (0.009) |
| *Asia* | -0.060*** | 0.002 | -0.006 |
| | (0.016) | (0.014) | (0.013) |
| *dummy 95* | -0.034*** | -0.037*** | -0.015** |
| | (0.007) | (0.007) | (0.006) |
| *dummy 97* | 0.103*** | 0.111*** | 0.127*** |
| | (0.007) | (0.005) | (0.006) |
| *dummy 98* | 0.112*** | 0.127*** | 0.127*** |
| | (0.007) | (0.006) | (0.006) |
| *differential int rates* | 0.002 | 0.001 | -0.003 |
| | (0.002) | (0.002) | (0.003) |
| *differential inflation* | 0.073*** | 0.081*** | 0.071 |
| | (0.014) | (0.013) | (0.010) |
| *Exchange rate regime* | 0.009 | -0.030*** | -0.014** |
| | (0.011) | (0.007) | (0.007) |
| J-Statistics | 204.18 | 207.85 | 213.68 |
| Instrument Rank | 144 | 177 | 177 |

Estimated model:$\rho_{ij,t} = \alpha_{ij} + \delta\rho_{ij,t} + \phi' x_{ij,t} + \gamma' z_t + \epsilon_{ij,t}$. White standard errors of estimates are in parentheses. All models include fixed individual effects. Individual effects not reported to save space.***, ** and * denote the significance of estimates at 10, 5 and 1 percent levels respectively. Column (1) use the measure of financial liberalization of Bekaert and Harvey (2003) while columns (2) and (3) use the two measures built by Edison and Warnock (2003).
The instrument space is composed by the lagged variables up to four lags.

allows us to claim that our main results are robust to specific econometric issues.

#### 4.2.3 Rolling correlations

Finally, a last robustness check involves an alternative measure of stock market co-movement. The use of realized moments at macroeconomic frequencies is rather new. Some of the underlying assumptions of the consistency of the realized moments that are fulfilled at the intra-daily frequencies might be less satisfied at the daily frequency. Therefore, in order to test for the robustness of the results with respect to the measure of the stock price co-movements, we use the usual coefficient of correlation between the pairs of daily returns over the calendar year as a measure of co-movement.

The results are reported in Table 5. We report the results using our three alternative measures. Equation (6) is estimated using GMM (with up to four lags for the instruments. On the whole, the results obtained with the traditional correlation are similar to those with the realized correlation. Financial liberalization is found to increase the co-movement of stock returns. The same applies to trade liberalization although the significance of the trade liberalization coefficient is lower with

Table 5: Impact on stock market rolling correlations: dynamic approach

| | Baseline model | Edison-Warnock 1 | Edison-Warnock 2 |
|---|---|---|---|
| *lagged correlation* | -0.039*** | -0.022*** | -0.014*** |
| | (0.014) | (0.006) | (0.006) |
| *financial liberalization* | 0.291*** | -0.197*** | -0.166*** |
| | (0.055) | (0.017) | (0.006) |
| *trade liberalization* | 0.006 | 0.001 | 0.025*** |
| | (0.009) | (0.003) | (0.003) |
| *trade intensity* | 0.063*** | 0.155*** | 0.296*** |
| | (0.003) | (0.012) | (0.022) |
| *Asia* | 0.080*** | 0.008 | -0.013*** |
| | (0.006) | (0.005) | (0.003) |
| *dummy 95* | -0.004*** | -0.046*** | -0.067*** |
| | (0.005) | (0.002) | (0.002) |
| *dummy 97* | 0.083*** | 0.117*** | 0.100*** |
| | (0.006) | (0.003) | (0.002) |
| *dummy 98* | -0.038*** | -0.027*** | -0.006*** |
| | (0.004) | (0.002) | (0.002) |
| *differential int rates* | 0.008*** | 0.007*** | 0.021*** |
| | (0.003) | (0.002) | (0.003) |
| *differential growth* | -0.004*** | -0.005*** | -0.007*** |
| | (0.001) | (0.001) | (0.002) |
| *differential inflation* | -0.006 | 0.003 | -0.051*** |
| | (0.009) | (0.005) | (0.005) |
| *Exchange rate regime* | -0.010 | -0.042*** | -0.067** |
| | (0.007) | (0.002) | (0.003) |
| J-Statistics | 256.90 | 274.09 | 263.74 |
| Instrument Rank | 234 | 273 | 273 |

Estimated model:$\rho_{ij,t} = \alpha_{ij} + \delta\rho_{ij,t} + \phi' x_{ij,t} + \gamma' z_t + \epsilon_{ij,t}$.. All models include fixed individual effects. Individual effects not reported to save space.***, ** and * denote the significance of estimates at 10, 5 and 1 percent levels respectively.
GMM estimation. The instrument space is composed by the lagged variables up to four lags.

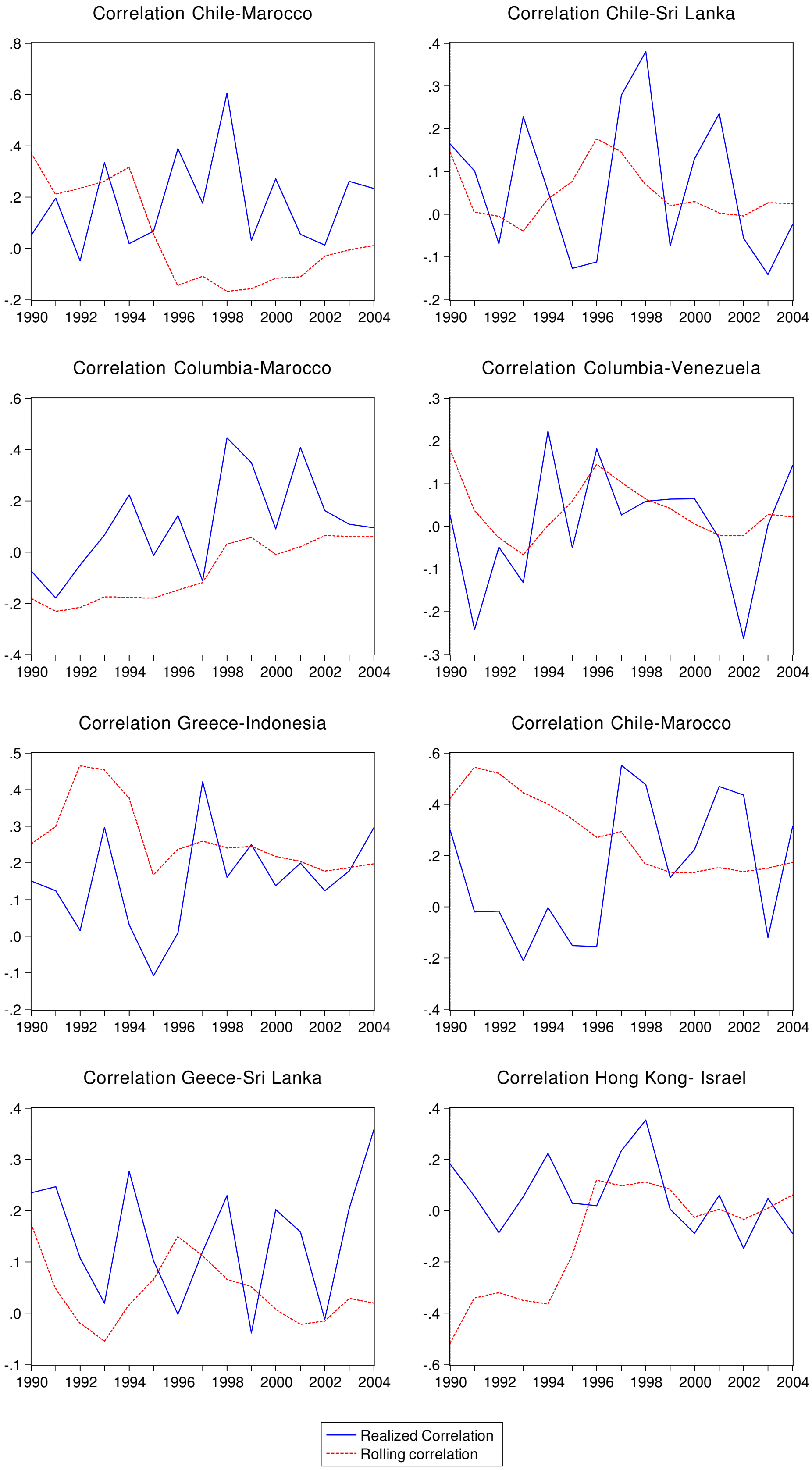
some measures of financial liberalization. This is partly due to the fact that the traditional correlation coefficients tend to smooth the variation of the co-movements, as reflected by Figure 1 for some country pairs. This in turn lowers the precision of some estimates. Trade intensity is also found to increase the co-movement of stock prices.

# 5 Conclusion

In this paper, we have studied the impact of trade and financial liberalization on the degree of stock market co-movement among emerging countries. We use a time-varying measure of stock returns correlations and combine cross-section and time series observations. This allows us to control for unobserved heterogeneity both across pairs of countries and over time, which allows to lower the probability of mispecification bias.

Our results provide strong support in favor of a positive impact of financial liberalization on stock market co-movement. This main finding is robust to alternative measures of capital openness, to the estimation technique and to the choice of the measure of stock price comovements.

**Figure 1: Realized vs Rolling Correlation**



The evidence provided in this paper is consistent with the idea that a decrease in asymmetric information due to the release of financial controls might favor cross-border financial flows. We also find that trade liberalization tends to increase the correlation between stock returns of the trading partners. Importantly, since we control for trade intensity, this effect goes through a signalling channel which tends to supplement the role played by the pure amount of trade. Our results show that policy reforms affecting both the real side and the financial system in emerging economies exert strong effects on the behavior of financial investors.

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